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Transparency of Retail Energy Pricing: Evidence from the US Natural Gas Industry

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We test and cannot reject the hypothesis that retail pricing of natural gas is transparent for commercial and residential customers served by regulated local distribution companies in the United States. The periods of adjustment to a wholesale price change are 1.54 months for the commercial price and 1.69 months for the residential price. These findings support the view that regulated local distribution companies quickly adjust retail prices to fully capture any change in the wholesale natural gas price. Copyright © 2013 John Wiley & Sons, Ltd.

1. INTRODUCTION

The retail energy prices paid by end users in the United States reflect, in part, the prices that their local distribution companies (LDCs) paid for that energy for resale in the competitive wholesale markets. Unregulated LDCs deliver propane, gasoline, and heating oil to end users, whereas regulated LDCs are typically responsible for end-use deliveries of electricity and natural gas. A regulated LDC sets its retail prices subject to a cost-of-service constraint and regulatory oversight (Bonbright *et al*., 1988), which means its retail prices are subject to *ex post* prudence review.

Those competitive energy wholesale markets are the result of market reforms and deregulation that sought to achieve via competition the heightened economic efficiency of the formerly vertically integrated energy industries (Newberry, 1999; MacAvoy, 2000; Sioshansi and Pfaffenberger, 2006). That wholesale price competition, in turn, is otherwise presumed to benefit end users via a 'transparent' pricing mechanism that can timely and fully transmit wholesale market price changes to retail prices. A well-known case in point is the real-time pricing of electricity (Schweppe *et al*., 1988; Stoft, 2002). In consequence, however, wholesale price volatility may require end users to manage their energy price risks (Pilipović, 1998; Eydeland and Wolyniec, 2003).

The purpose of this paper is to present the results of our analysis of a decade's worth of retail pricing data for one particular energy market—the natural gas

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market in the United States. The intent of the analysis is to determine whether retail natural gas prices have indeed been as transparent as the market reformers intended them to be. Our analysis is timely and relevant because of the currently low wholesale natural gas prices caused by a slowly recovering economy and rapidly rising shale-gas output. It is based on a direct test that we proffer for the transparent retail pricing of energy. We implement the test using US monthly natural gas market data for the period from 2001 to 2010.

We use the most recent 10-year period to implement our test for the following reasons. First, natural gas price data in the period reflects the current market environment, which emerged after the last major order (Order 637) issued in 2000 by the Federal Energy Regulatory Commission (http://www.naturalgas.org/regulation/market.asp). Second, the period contains both high-price and low-price months required for precisely identifying the wholesale–retail price relationship that we postulate later in the text. Finally, it provides a sufficiently large sample size (120 observations) so that our results are not likely to be subject to small-sample bias.

The transparent pricing issue is important from management's perspective because a failure to fully and quickly pass through to retail customers any wholesale price changes can bring downside risks. When wholesale prices go down, slow changes in retail prices can be seen as management's attempt to exploit the LDC's customers. By contrast, when wholesale prices go up, the LDC's failure to be compensated for its increased cost translates, in effect, into a loan to those customers. As those uncompensated loans increase, the risk of disallowance by its regulator can occur (Ryan and Lieberman, 2012). Although regulation has been relaxed in this industry, various regulatory or government pressures and rules persist encouraging commodity costs to be passed through with minimal markup or discount. Hence, we hypothesize that the LDC will quickly adjust its retail prices to fully capture any changes in the wholesale prices that it pays. But does this hypothesis have the support of real world data?

We choose natural gas to answer this question for the following reasons. First, natural gas market reform has led to wholesale market competition throughout North America (De Vany and Walls, 1995; MacAvoy, 2000; Mariner-Volpe and Trapmann, 2003; Walls, 2008). Second, notwithstanding market power in some wholesale natural gas market hubs (Murry and Zhu, 2008) and limited arbitrage due to pipeline capacity shortages (Brown and Yücel, 2008), ample evidence has shown that the trading hubs in the USA are highly competitive and tightly integrated (De Vany and Walls, 1993, 1994; Walls, 1994; Doane and Spulber, 1994; NEB, 1995; King and Cuc, 1996; Serlatis, 1997; Kleit, 1998; Cuddington and Wang, 2006; Park *et al.*, 2007; Gebre-Mariam, 2011). Finally, there is evidence that wholesale and retail markets are co-integrated in the statistical sense (Mohammadi, 2011), the implication of which is that 'even the smallest volume natural gas customers—residential consumers—have felt the benefits of the industry restructuring' (Arano and Velikova, 2009, p. 129).

We recognize that Mohammadi (2011) has comprehensively studied bidirectional Granger causality in the natural gas industry and concluded that in the long run 'demand shocks (are) the primary determinants of natural gas prices' (p. 227). The focus of our paper, however, is on the important dual managerial issues of price transparency and of management's short-term response to changes in its commodity procurement costs in the process of delivering natural gas to its retail customers. Our particular concern is to determine if a regulated LDC's management would quickly and fully transmit a wholesale price change to its retail prices so as to preempt an unwelcome *ex post* prudence review.

The paper makes three principal contributions. First, our test of transparent pricing is based on the cost-of-service standard commonly used by a regulated LDC when setting its retail prices (Bonbright *et al.*, 1988), thus complementing studies based on co-integration and Granger causality (e.g., Arano and Velikova, 2009; Mohammadi, 2011). Second, we find that commercial and residential retail prices closely follow the cost-of-service standard. Finally, we show that commercial and residential retail prices adjust quickly in response to changes in the wholesale price. Thus, we conclude that retail pricing of natural gas by regulated LDCs in the USA is indeed transparent.

2. THE MODEL AND ESTIMATION PROCEDURE

2.1. Model Specification

Consider a regulated LDC whose per unit noncommodity cost ($/million British thermal units in month t) is denoted C_t. The magnitude of C_t, which reflects the LDC's noncommodity embedded costs for customer service and delivery (Harunuzzaman and Koundinya, 2000),

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may be subject to price-cap regulation aimed at improving the LDC's productivity (Liston, 1993; Laffont and Tirole, 1993; Schmidt, 2000; Vogelsang, 2002).

Let P_t denote the wholesale natural gas price in month t. Further let R_t^* denote the unobservable long-run equilibrium to which the retail price will gravitate from the observable short-run price of R_t. We say R_t^* is transparent when it is strictly cost based and equal to the arithmetic sum of its two cost components:

$$R_t^* = C_t + P_t. \tag{1}$$

We consider R_t^* as an equilibrium price, in the sense that, once reached, it will rest undisturbed unless there are nonoffsetting changes in its two underlying cost components.

Retail price transparency implies that retail buyers can confidently expect that, on the one hand, any reductions in wholesale prices will *ultimately* be reflected in the prices that they pay in the retail market, whereas on the other hand they will understand that any price increases they are asked to absorb in the short run are in response to corresponding increases in the competitively determined wholesale prices that the LDCs pay. Moreover, that confidence should be shored up by (i) the Federal Energy Regulatory Commission's established procedures 'for the collection and publication of volume and price information' for natural gas market transactions on a daily and monthly basis as well as (ii) 'real-time transparency for futures, options, and a large (and growing) number of swaps' in the financial markets (Albrecht, 2009, pp. 9–10).

To be sure, there are reasons why the LDC's retail pricing may *not* be transparent. First, full transparency requires immediate and complete transmission of the changes in the LDC's cost components to the retail price, which may not occur because of the LDC's infrequent retail price adjustments. Second, even if retail price adjustments can occur monthly, the LDC only bills its customers for their consumption in the prior month. Real-time pricing has yet to be implemented in the natural gas retail market. Third, the LDC may follow a regulatory mandate of rate stability, only passing through a portion of the change in the wholesale price to the retail price. Finally, the LDC may respond asymmetrically to wholesale price changes. In particular, it may quickly raise the retail price when faced with rising wholesale prices while slowly lowering it when wholesale prices fall. Hedging programs to reduce retail price volatility may further complicate these relationships by (i) creating differences between the prices actually paid by the LDC and the commodity prices set in the wholesale market and (ii) delaying the exposure of the LDC to changes in the wholesale price. But these considerations beg the question of whether the economically meaningful and easy to understand Equation (1) does in fact describe the LDC's retail pricing from a long-run perspective, one that must account for the unobservable R_t^*.

Our approach to the problem is to conduct an econometric analysis, described later in the text, that relies on the observable short-run monthly prices. Those price series, however, are almost surely nonstationary, which can result in a spurious regression problem (Davidson and MacKinnon, 1993, pp. 669–673). In anticipation of that problem, as our launch point, we take the first differences ($\Delta X_t = X_t - X_{t\text{-}1}$) for each of the three time series to obtain:

$$\Delta R_t^* = \Delta C_t + \Delta P_t. \tag{2}$$

Then, to finesse the problem of the unobservable long-run prices, first let R_t' denote the retail price that LDC management has targeted for the long run for month t in response to its noncommodity costs and the wholesale price paid in the natural gas market. We then write the counterpart of Equation (2) as follows:

$$\Delta R_t' = \varphi \Delta C_t + \theta \Delta P_t. \tag{3}$$

Equation (3) allows $\Delta R_t'$ to differ from ΔR_t^* via the parameter φ for ΔC_t and θ for ΔP_t. For example, if $\varphi = 1.1$, a \$1 change in the noncommodity cost leads to a \$1.10 change in the targeted long-run retail price. When $\varphi = \theta = 1$, $\Delta R_t' = \Delta R_t^*$.

Second, because of institutional factors and the market frictions that abound in the real world, management may be unable to immediately implement its targeted long-run prices in the short run. Thus, we never actually observe those prices. Rather, what we observe is R_t, the *actual* retail price charged in month t, which may differ from R_t'.

Let ΔR_t denote the dollar change in the LDC's actual retail price, where ΔR_t may differ from $\Delta R_t'$. Hence, we assume a partial adjustment process such that:

$$\Delta R_{t\cdot} - \Delta R_{t-1} = (1 - \lambda)\left(\Delta R_t' - \Delta R_{t-1}\right), \tag{4}$$

where $0 < \lambda \leq 1$ is the coefficient of adjustment.

In tandem, Equations (3) and (4) yield the following specification:

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$$\Delta R_t = (1-\lambda)\varphi\Delta C_t + (1-\lambda)\theta\Delta P_t + \lambda\Delta R_{t-1}. \quad (5)$$

We cannot estimate the parameters of Equation (5), however, because we do not have actual data on ΔC_t. To circumvent *this* unobservable data problem, we assume the following linear approximation that has a time trend and an approximation error ε_t:[1]

$$(1-\lambda)\varphi\Delta C_t = \alpha + \beta t + \varepsilon_t. \quad (6)$$

With this assumption, the regression to be estimated is:

$$\Delta R_t = \alpha + \beta t + (1-\lambda)\theta\Delta P_t + \lambda\Delta R_{t-1} + \varepsilon_t. \quad (7)$$

Equation (7) is a nonlinear regression equation whose coefficients to be estimated are (α, β, θ, λ).[2]

2.2. The Estimation Procedure and Hypotheses

Our estimation procedure has four steps which we apply to each of the two end-user classes: commercial and residential. We focus on a regulated LDC's commercial customers (e.g., small offices, schools, restaurants, and retail stores) and residential customers (e.g., single family homes, condos, and apartments) for two reasons. First, they tend to be small in size and are the LDC's captive customers, likely protected by a regulator with the threat of *ex post* prudence review and cost disallowance. Second, industrial users of natural gas (e.g., manufacturing plants and power generation stations) are much larger in size and can take advantage of open access to pipelines to buy directly from the wholesale market without using the service of regulated LDCs (http://www.naturalgas.org/regulation/market.asp).

The steps are as follows:

- Step 1: Apply the Phillips–Peron unit root test (Phillips and Perron, 1988) to the two retail price data series as well as to two wholesale price data series mentioned later in the text.
- Step 2: Estimate Equation (7) for each end-user class to determine whether the residuals are serially correlated.
- Step 3: Jointly estimate Equation (7) for both end-user classes as a pair of nonlinear seemingly unrelated regressions (Gallant, 1987). We use the seemingly unrelated regressions approach to take into account the likely possibility that the retail prices in both end-user classes are impacted by correlated errors due to common random factors (e.g., weather conditions).
- Step 4: Perform a Wald test of the two null hypotheses that go to the heart of the price transparency issue.

The first null hypothesis is H_0: $\theta = 1$. Under this hypothesis, Equation (3) shows that a \$1 change in the wholesale price translates into a \$1 change in the retail price. The translation, however, may not be immediate simply because a regulated LDC bills its customers monthly for their consumption in the prior month. Thus, we expect that the period of adjustment of $1/(1-\lambda)$ month is likely to be longer than one month.

The second null hypothesis is H_0: $\theta = 1$ and $\lambda = 0$. Under this hypothesis, Equation (7) affirms that management establishes a transparent retail pricing policy such that a change in the wholesale price is immediately felt by its customers.

3. THE DATA SAMPLE

We apply this test to national level monthly natural gas prices for the United States using publicly available data from the US Department of Energy's Energy Information Agency (http://www.eia.gov/dnav/ng/ng_pri_sum_dcu_nus_m.htm) for the period from 2001 to 2010 resulting in 120 monthly observations. Our data sample has two wholesale price series: wellhead and citygate. Wellhead prices are what natural gas producers receive at the wellhead. They include 'all costs prior to shipment from the lease, including gathering and compression costs, in addition to State production, severance, and similar charges' (http://205.254.135.7/dnav/ng/TblDefs/ng_prod_whv_tbldef2.asp). Citygate prices are what the LDCs pay at the point at which they obtain the gas from the pipeline (e.g., Brown and Yücel, 1993, p. 41; Mohammadi, 2011, p. 227). Figure 1 shows that the retail prices tend to track the wellhead price. Figure 2 shows that they move closely with the citygate price.

Panel A of Table 1 shows that the average wellhead price of \$5.31/thousand cubic feet (Mcf) is the lowest, below the average citygate price of \$6.86/Mcf, the average commercial price of \$9.90/Mcf, and the average residential price of \$12.41/Mcf. The standard deviations indicate that the wellhead prices are the least volatile, followed by the citygate, commercial, and residential prices. The range defined by the difference between the maximum and minimum price is the smallest for the wellhead price and the largest for the residential price.

Panel B reports the correlation coefficients, showing that the two wholesale prices are highly correlated ($r = 0.95$), the citygate price is highly correlated ($r = 0.96$) with the commercial price, and less so

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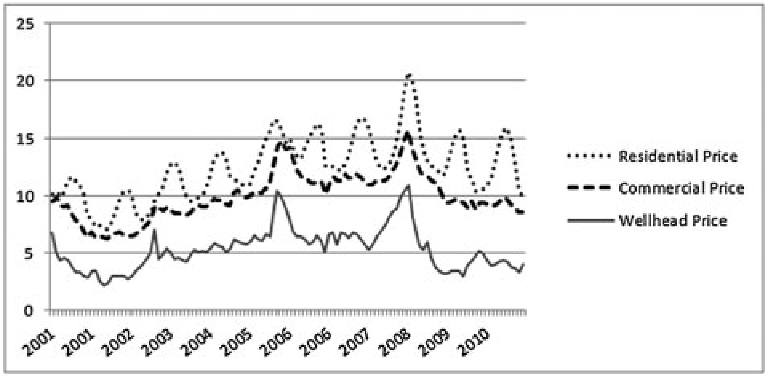

Figure 1. Monthly natural gas retail prices (residential and commercial) versus wholesale prices (wellhead) for the period of January 2001–December 2010 ($/thousand cubic feet [Mcf]).

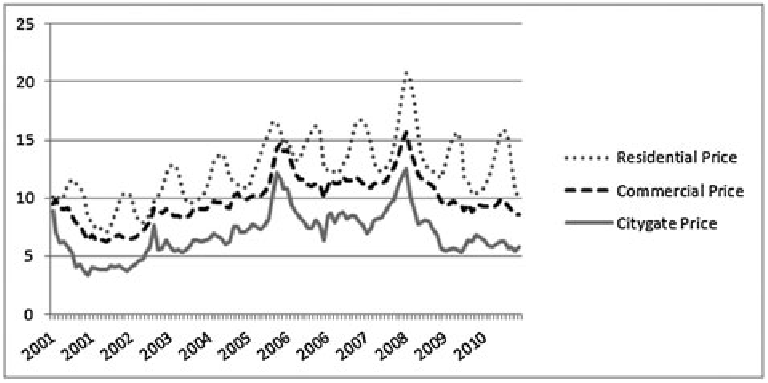

Figure 2. Monthly natural gas retail prices (residential and commercial) versus wholesale prices (citygate) for the period of January 2001–December 2010 ($/thousand cubic feet [Mcf]).

($r = 0.71$) with the residential price. Finally, the two retail prices are highly correlated ($r = 0.82$).

Panel C reports the descriptive statistics for the first difference data, showing that the price difference series have zero means, are less volatile than the price-level series and have narrower ranges than the price-level series. Panel D reports correlation coefficients that are generally lower than those in Panel B. Again, however, the wellhead price differences and citygate price differences are highly correlated ($r = 0.9$) as are the citygate price differences and the commercial price differences ($r = 0.82$).

4. THE RESULTS

Our regression results are based on a sample of $N = 118$ observations because one observation is 'lost' in the lagging process, and a second meets the same fate as a result of first differencing.

4.1. The Step 1 Results

In Step 1, we conduct Phillips–Perron unit root tests for the retail price data series as well as for the wholesale price data series. Panel A of Table 2 shows that the null hypothesis of a nonstationary data series cannot be rejected at the 5% level.

The nonstationarity of the retail price series suggests that we employ the first difference format that corresponds to Equation (7). Panel B of Table 2 shows that the null hypothesis of nonstationarity can be rejected at the 1% level, implying that the first difference data should be used for our subsequent analysis.

4.2. The Step 2 Results

Step 2 requires that we regress ΔR_t against ΔP_t, $\Delta R_{t\text{-}1}$, and a time trend to explore patterns in the residuals. The estimated coefficient for the time trend variable was statistically insignificant for both customer classes ($p > 0.75$), further confirming that the first difference

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Table 1. Monthly Natural Gas Prices for January 2001–December 2010

Price level ($/Mcf)	Mean	Standard deviation	Minimum	Maximum
Panel A: Descriptive statistics based on price-level data				
Wellhead	5.31	1.85	2.19	10.79
Citygate	6.86	1.95	3.37	12.48
Commercial	9.90	2.04	6.28	15.64
Residential	12.41	2.78	7.10	20.77
Panel B: Correlation coefficients based on price-level data				
Price level ($/Mcf)	Wellhead	Citygate	Commercial	Residential
Wellhead	1.00	0.95	0.88	0.63
Citygate	0.95	1.00	0.96	0.71
Commercial	0.88	0.96	1.00	0.82
Residential	0.63	0.71	0.82	1.00
Panel C: Descriptive statistics based on price difference data				
First difference ($/Mcf)	Mean	Standard deviation	Minimum	Maximum
Wellhead	−0.02	0.71	−2.58	2.47
Citygate	−0.03	0.70	−2.28	2.06
Commercial	−0.01	0.52	−1.45	1.70
Residential	0.00	1.03	−3.20	2.58
Panel D: Correlation coefficients based on price difference data				
First difference ($/Mcf)	Wellhead	Citygate	Commercial	Residential
Wellhead	1.00	0.90	0.69	0.25
Citygate	0.90	1.00	0.82	0.29
Commercial	0.69	0.82	1.00	0.52
Residential	0.25	0.29	0.52	1.00

Mcf, thousand cubic feet.

Table 2. Phillips–Perron Unit Root Test Results: τ Statistics with Two Lags and p-Values in ( )

Type	Wellhead price	Citygate price	Commercial price	Residential price
Panel A: Results based on price-level data that have a nonzero mean				
Single mean	−2.359 (0.156)	−2.271 (0.183)	−1.652 (0.453)	−2.839 (0.057)
Single mean with trend	2.433 (0.361)	−2.579 (0.291)	−1.670 (0.759)	−3.121 (0.107)
Panel B: Results based on price difference data that have a zero mean				
Type	Wellhead price	Citygate price	Commercial price	Residential price
Zero mean	−9.480 (0.001)	−9.436 (0.001)	−7.507 (0.001)	−5.526 (0.001)

time series are stationary. In addition, Table 3 shows that the regression residuals are serially correlated. Hence, we use a first-order autoregressive (AR(1)) estimation procedure in Step 3.

4.3. The Step 3 Results

In Step 3, we obtain estimates of β for each customer class. With $p > 0.63$, the two estimates are highly insignificant. Hence, there is no significant trend over time in the noncommodity portion of retail natural gas prices. Indeed, recent analyses from the US Government Accounting Office report very little change in the average level of noncommodity local distribution charges since 2001 (U.S. GAO, 2006, p. 12). Further, the two estimates of α have $p > 0.65$. Finally, the Wald statistic of 0.26 ($p = 0.99$) suggests that we cannot reject the null hypothesis of $\alpha = \beta = 0$ even at the 10% level for both customer classes. We take advantage of these findings to remove $\alpha + \beta t$ from the model in repeating Step 3 of the estimation process.

Table 3. First-order Autoregressive Parameter Estimates for Regression Residuals with p-Values in ( )

Wholesale price	Commercial price	Residential price
Citygate	−0.477 (<001)	0.345 (0.047)
Wellhead	−0.457 (<001)	0.128 (0.435)

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Table 4. Seemingly Unrelated Regression Results with *p*-Values in ( )

Estimate	Commercial price	Residential price
Panel A: Wholesale price = Citygate		
Adjusted R^2	0.856	0.557
λ	0.349 (<001)	0.409 (0.004)
θ	0.980 (<001)	1.106 (0.002)
AR(1) parameter	−0.363 (<001)	0.359 (0.021)
Wald statistic for testing H_0: $\theta = 1$	0.160 (0.691)	0.090 (0.761)
Wald statistic for jointly testing H_0: $\theta = 1$ and $\lambda = 0$	260.53 (<001)	36.44 (<001)
Panel B: Wholesale price = Wellhead		
Estimate	Commercial price	Residential price
Adjusted R^2	0.680	0.483
λ	0.433 (<001)	0.502 (0.004)
θ	0.941 (<001)	1.000 (0.022)
AR(1) parameter	−0.243 (0.005)	0.294 (0.132)
Wald statistic for testing H_0: $\theta = 1$	0.390 (0.534)	0.000 (0.999)
Wald statistic for jointly testing H_0: $\theta = 1$ and $\lambda = 0$	188.64 (<001)	43.94 (<0001)

AR(1), first-order autoregressive.

Table 4 presents the regression results from Step 3 wherein we estimate the parameters of Equation (7) with both α and β set equal to zero. All of the remaining parameter estimates are virtually identical to those found for the more general case of $\alpha \neq 0$ and $\beta \neq 0$.

The estimates in Panel A of Table 4 are based on the wholesale price being the citygate price. The commercial price difference regression's adjusted R^2 is 0.856, which is higher than the residential price difference regression's adjusted R^2 of 0.557. This is to be expected because a regulator tends to shield residential customers from large wholesale price spikes more than it does commercial customers. All estimates for λ and θ, as well as for the two AR(1) parameter estimates, are significantly different from zero ($p \leq 0.004$), aside from the AR(1) estimate associated with the residential sector ($p = 0.021$).

The AR(1) parameter estimates suggest moderate negative autocorrelation for commercial retail pricing, signifying the oscillating effect of a past random error. There is moderate positive autocorrelation for residential retail pricing. In either event, we are able to reject the hypothesis ($p < 0.01$) that $|\text{AR (1) Coefficient}| = 1$. The implication is that the random error term has a finite variance, which obviates our concerns as to whether we have a spurious regression problem.

As an additional check, we use wellhead prices to re-estimate Equation (7). Panel B of Table 4 reports these coefficient estimates, which are very similar to those in Panel A. This is unsurprising given the historically high correlation of the wholesale prices ($r = 0.95$) in our sample and indeed more generally in the co-integration of wellhead and citygate prices (Mohammadi, 2011, p. 230).

4.4. The Step 4 Results and Hypothesis Tests

The bottom rows of Panel A of Table 4 present the results of the Wald statistics for testing the two null hypotheses related to retail pricing transparency when the wholesale price is the citygate price. As seen from the table, we fail to reject the null hypothesis that $\theta = 1$, lending support to the hypothesis that the retail pricing rule for commercial and residential customer classes is consistent with exact cost-based pricing. We soundly reject, however, the *joint* hypothesis that $\theta = 1$ and $\lambda = 0$, which implies that a regulated LDC's management does not have a transparent pricing rule that results in a wholesale price change that is *immediately* felt by end users. This is not surprising because the LDC's commercial and residential customers are typically billed monthly for their consumption in the prior month. Further, the use of hedging programs by utilities may also create some lag in the adjustment of retail prices to changes in wholesale prices. Nonetheless, the periods of adjustment are estimated to be quite short, 1.54 months for commercial customers and 1.69 months for residential customers.

The bottom rows of Panel B of Table 4 report the Wald statistics that show replacing the citygate price with the wellhead price as the wholesale price does not alter our finding of retail pricing transparency. The periods of adjustment, however, are now 1.76 months for commercial customers and 2.00 months for residential customers.

5. CONCLUSION

Carter and Curry (2010) document that consumers value price transparency as opposed to opaque pricing and might in fact be willing to pay more for a transparently priced product (p. 768). In this paper, we have presented a generally applicable four step econometrically based process for verifying long-run price transparency from a wholesale market to the various retail markets that depend upon it. Most critically, we have demonstrated the practicality of that process by applying it to analyze price transparency in an economically important and vital segment of the US economy: notably, the natural gas industry for the first decade of the new millennium.

Our initial conclusion was that wholesale and retail prices are nonstationary. As a result, we employed a first difference format in the natural gas application. We further modify our procedure to take account for serially correlated disturbances by adopting AR(1) modeling.

Our results affirm retail price transparency for commercial and residential retail energy consumers in the restructured natural gas market in the United States. The prices of natural gas sold to commercial and residential retail energy consumers quickly adjust to wholesale price changes with adjustment periods of 1.54 and 1.69 months. Although our analysis of national data cannot confirm whether these results would hold for all LDCs and in all regulatory environments, these results suggest a general tendency for LDC management to quickly adjust its retail prices to fully capture any changes in the wholesale prices, thereby preempting the downside risks of an *ex post* prudence review by its regulator.

Finally, application of the procedures presented here could be applied to a specific LDC to explore its price transparency. A demonstration of transparency provides evidence of a properly functioning market and responsible utility behavior. Utility-specific results could then be compared with an industry average benchmark to explore a particular utility's relative transparency and speed of adjustment in retail prices.

NOTES

1. We do not know, *a priori*, if there is a general inflation trend of C_t, the monthly unobserved component of the retail price. Hence, we use a linear approximation of ΔC_t, which results from a second-order quadratic approximation of an unknown function for C_t. If the estimates for α and β are statistically insignificant, we can exclude these coefficients from the regression analysis. Section 4.3 reports that the estimates for α and β are highly insignificant, and their exclusion does not cause concerns of under-specification that may bias our regression results in Table 4 later in the text.
2. Although we have verified an order-$I(1)$ co-integration relationship between the two retail prices and the wholesale prices of natural gas in *our* database, for several reasons, we eschewed the opportunity to incorporate an error correction mechanism (ECM) into our regression analysis. First, an ECM does not account for the situations depicted in Equations (3) and (4), and thus its use will lead to misspecification of the model. Second, our focus is to investigate the effect of a wholesale price change on the retail price change rather than the long-run relationship between two price-level series. Finally, the commodity cost adjustment clause in regulated retail ratemaking firmly establishes that the direction of causality runs from a monthly wholesale price change to a retail price change, thus obviating the use of an ECM as a means for examining the causal relationships.

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